GENDER PREDICTION FROM BODY MEASUREMENTS OF TWO SUBSPECIES OF SANDHILL CRANES

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Abstract: Linear discriminant functions estimated from leg length, bill length and body weight measurements of known subspecies and known gender (KS-KG) Florida (FSH) and greater sandhill cranes (GSH) (*Grus canadensis pratensis* and *G. c. tabida*, respectively) were used to predict gender of known subspecies and unknown gender (KS-UG) individuals. Mean body measurements were larger among males than among females of either subspecies, but the difference in mean bill length between genders was larger among GSH than among FSH cranes. Gender misclassification was less frequent among GSH than among FSH cranes. Of birds whose measurements fell outside of the 80% prediction region surrounding either gender measurement mean, misclassification rates were ≤6%.

Key Words: body measurements, Florida, gender prediction, *Grus canadensis*, linear discriminant function, sandhill cranes, subspecies

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Fundamental biological features of a species or population (i.e., survival rate, productivity, and age of first breeding) may vary with gender. Gender is not apparent in monochromatic sandhill cranes, even when birds are in hand. Correct identification of gender is vital to understanding gender effects on demographic characteristics.

Gender in cranes could be determined using fecal steroids analysis (Czekala and Lasley 1977), chromosomal analysis of blood (Hungerford et al. 1966), feather karyotyping (Rasch and Kurtin 1976, Van Tuinen and Valentine 1987), or laparoscopy (Greenwood 1983). However, these techniques are invasive, expensive when applied to a large number of individuals, or effective only under certain conditions (molting, developed gonads, etc.). Gender determination by cloacal examination (Blackman 1971) is marginally successful (66% correct prediction) in sandhill cranes (Tacha and Lewis 1979).

Field observation of sandhill crane behavior provides another method for gender determination in sandhill cranes. The unison call, a gender-specific behavior (Archibald 1975), is most often displayed by a dominant pair on or near a defended territory, or by birds involved in agonistic encounters (Tacha 1988). The opportunity for observing unison calling by an individual bird can depend on age, status, location, or time of year. The infrequency of the call by a specific individual might preclude observing that individual engaged in the unison call for several years. Precopulation/copulation posture is usually a reliable indicator of gender, but reverse mounting does occur in sandhill cranes (S. A. Nesbitt, pers. observ.; G. W. Archibald, pers. commun.). Gender may be inferred from non-reproductive, gender-specific behavior (Tacha et al. 1987, Tacha 1988).

Males are usually larger than females within each of the 2 subspecies of sandhill cranes that occur in Florida (Walkinshaw 1949). If gender could be predicted with acceptable accuracy from measurements, then a known-gender sample population could be developed as part of a routine capture and marking operation. We tested gender dimorphism within subspecies and produced a mensural model to predict gender when subspecies is known but gender is not.

METHODS

Sandhill cranes caught in Florida, 1973–89, were individually color-marked (Nesbitt et al., in press). Those seen in Florida year-round were assigned to the FSH subspecies. Those seen outside Florida or seen in Florida only between 1 November and 31 March (the overwintering period for GSH [Nesbitt et al., in press]) in ≥2 wintering periods were assigned to the GSH subspecies. To develop a KS-KG sample subset, a bird's gender was determined from unison call posture or feather karyotyping. Other known-subspecies birds were assigned to a KS-UG subset.

Bill length (*B*; mm, posterior nare to tip), leg length

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(L; mm, tarsus plus exposed tibia), and weight (W; kg/100) were determined for all captured birds ≥18 months old. For the KS-KG subset, probability plots (Chambers et al. 1983:193-212) were used to assess normality of the marginal distribution of each measurement variable by subspecies-gender combination. Sample covariance matrices of W, B, and L were calculated for each subspecies-gender grouping. These matrices were tested for equality (Box 1949), then pooled to form a common sample covariance matrix used in all further analyses. Multivariate analysis of variance (Johnson and Wichern 1982:268-277) was used to detect differences in the W, B, and L measurement means among the factors gender, subspecies, and the gender by subspecies interaction.

For classification of gender given W, B, and L, linear discriminant functions (Johnson and Wichern 1982:466) were constructed for each subspecies (y_{FSH} and y_{GSH}) assuming equal prior assignment probabilities and misclassification costs. For facilitating direct comparison among coefficients in y_{FSH} and y_{GSH}, measurements were standardized to a common mean and variance prior to discriminant function estimation. To predict gender membership, we calculated values of y_{FSH} or y_{GSH} for each W, B, and L measurement triplet and classified the bird as female if y_{FSH} (or y_{GSH}) ≥ 0 and as male otherwise.

Because small sample sizes precluded our dividing the data into training and validation sets, we followed Lachenbruch's (1975) jackknifing procedure for unbiased estimation of misclassification rates. Each point in the KS-KG subset was held out (jackknifed) in turn, and the discriminant function was estimated from the remaining points. The excluded point was then classified, and the success of its classification was recorded. After all points were jackknifed, the misclassification rate was estimated as total number of misclassifications divided by sample size.

We attempted to reduce misclassification rates by developing a rule to allow us to exclude from classification those measurements falling in a "too-close-to-call" (TCTC) category. The risk of gender misclassification is small if a measurement triplet is relatively close to a gender mean. "Closeness" to a mean may be expressed in terms of the size of the smallest surrounding $(1 - \alpha) \cdot 100\%$ prediction region (the multivariate analog of the prediction interval) that contains the measurement. Thus, if containment of a measurement requires a large prediction region around a mean, then gender misclassification likelihood is small if the point is also contained in a small prediction region around the other mean. Conversely, gender classification is risky if point containment requires large regions around both means. For every jackknifed observation within each subspecies, we recorded the size of the smallest prediction region around each gender mean that would contain the observation. We plotted the greater of these 2 distances against the jackknifed discriminant function score for the observation. We selected an arbitrary distance cut-off value that most effectively separated correctly from incorrectly classified birds. Birds whose distances fell below the cut-off value were considered TCTC. We estimated the cost of the rule by dividing the number of correctly classified birds in the TCTC group by the sum of the correctly

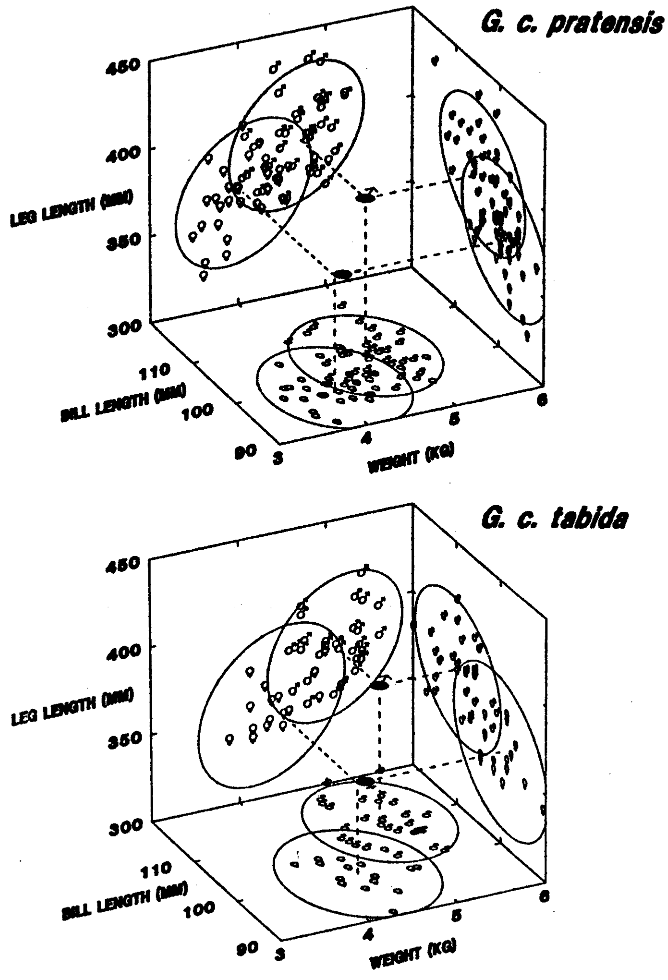

Fig. 1. Three-dimensional scatterplots, by subspecies and gender, of mean weight (kg), bill length (mm), and leg length (mm) (closed symbols) of 114 sandhill cranes ≥18 months old caught in Florida, 1973–89. Two-dimensional shadows of individual data points and of 80% prediction regions are displayed on the bivariate faces of the plot. Relative positions of the shadows of the prediction regions indicate that *tabida* cranes were heavier, longer-billed, and shorter-legged than *pratensis* cranes, and that males of either subspecies were heavier, longer-billed, and longer-legged than females. Gender disparity in bill length was greater in *tabida* than in *pratensis* cranes.

Table 1. Weight (*W*, kg), bill length (*B*, mm), and leg length (*L*, mm) means and standard errors, by gender and subspecies, of 114 sandhill cranes ≥18 months old caught in Florida, 1973–89. Unweighted mean differences,[a] difference standard errors, and results of tests of equality to zero are provided in the table margins.

		Gender				Mean difference between genders	
		Female		Male			
Subspecies	Variable	$\bar{x}$	SE	$\bar{x}$	SE	$\bar{x}$	SE
	W	4.066	0.058	4.669	0.058		
G. c. pratensis	*B*	98.8	0.73	103.9	0.73		
	L	363.6	3.1	390.1	3.1		
	W	4.374	0.084	5.101	0.062	−0.665[b]	0.066
G. c. tabida	*B*	100.2	1.06	109.7	0.78	−7.3[b]	0.84
	L	351.8	4.5	377.8	3.3	−26.2[b]	3.6
Mean difference	*W*			−0.370[b]	0.066	0.062	0.066
between	*B*			−3.6[b]	0.84	2.2[b]	0.84
subspecies	*L*			12.1[b]	3.6	−0.3	3.6

[a] Mean diff. between genders = $0.5[(\bar{x}_{P,F} - \bar{x}_{P,M}) + (\bar{x}_{T,F} - \bar{x}_{T,M})]$, mean diff. between subspecies = $0.5[(\bar{x}_{P,F} - \bar{x}_{T,F}) + (\bar{x}_{P,M} - \bar{x}_{T,M})]$, interaction = $0.5[(\bar{x}_{P,F} - \bar{x}_{P,M}) - (\bar{x}_{T,F} - \bar{x}_{T,M})]$, where P = *G. c. pratensis*, T = *G. c. tabida*, F = female, and M = male.

[b] Test of zero difference rejected at 0.01 level.

classified birds.

TCTC membership was determined for every observation in the KS-UG subset. Of those not TCTC, gender was predicted using the functions y_{FSH} or y_{GSH}.

RESULTS AND DISCUSSION

The KS-KG calibration sample subset consisted of 68 FSH (34 females, 34 males) and 46 GSH (16 females, 30 males). We detected no significant departures from normality in the univariate normal probability plots. Bivariate scatter plots also revealed no obvious departures from bivariate normality (Fig. 1). A test of equality of the 4 subspecies-gender covariance matrices was not rejected (Box's M = 18.7, 18 df, P = 0.412).

For either gender, GSH were heavier, shorter-legged, and longer-billed than FSH (F = 27.5; 3, 108 df; P < 0.001). For either subspecies, males were heavier, longer-legged, and longer-billed than females (F = 60.1; 3, 108 df; P < 0.001). All 6 univariate tests of the subspecies and gender effects were highly significant (Table 1). The difference in gender means varied by subspecies (F = 2.79; 3, 108 df; P = 0.004); specifically, in bill length. Whereas bills were longer among males than females of either subspecies, the difference was greater in GSH than FSH (t = 2.67, 110 df, P = 0.009, Table 1). Relative position of bivariate prediction ellipses (Fig. 1) confirmed these relationships.

Negative discriminant function coefficients indicated that birds of smaller dimensions were more likely classified female (Table 2). Gender means were highly distinct for either subspecies, but distinction was greatest in GSH (Table 2). In GSH, *B* was a more discriminating measurement for gender classification than in FSH (Table 2, Fig. 1). For either subspecies, *W* was a better discriminator than *L* (Table 2).

Jackknife application of the discriminant function to the KS-KG data produced 11 (16%) gender misclassifications in FSH and 3 (7%) in GSH. A point lying outside of 1 or both 80% prediction regions surrounding the gender means appeared less likely to be misclassified than a point contained within both regions (Fig. 2). By declaring points within both 80% prediction regions TCTC, misclassification rates improved to 3 of 51 (6%) for FSH and 0 of 41 (0%) for GSH at a cost of 9 of 48 (19%) otherwise correctly classified FSH and 2 of 41 (5%) GSH. Of 18 KS-UG FSH, gender was predicted for 14 (78%, 5 M, 9 F) that were not TCTC. Of 233 GSH, gender was predicted for 200 (86%, 102 M, 98 F).

A mensural-driven estimator of gender would be inappropriate in small-sample situations when, for example, associations of specific behaviors with gender are made. In these situations, gender should be determined by a more error-free method; biochemical methods have

Table 2. Estimated gender discriminant function coefficients for standardized values of weight (W_s), bill length (B_s), and leg length (L_s), by coefficient type[a] and subspecies, of 114 known-gender sandhill cranes ≥18 months old caught in Florida, 1973–89. Sample squared distance between gender population means (D^2) and P for the hypothesis $D^2 = 0$ are provided.

Type	Subspecies	D^2	P	Coefficient β_W	β_B	β_L
unscaled	*G. c. pratensis*	5.18	<0.001	−2.328	−1.433	−1.071
	G. c. tabida	9.69	<0.001	−2.999	−2.982	−0.652
scaled	*G. c. pratensis*			−0.793	−0.488	−0.365
	G. c. tabida			−0.701	−0.697	−0.152

[a] Substitute unscaled coefficients in the following model to predict gender: Female: $\beta_0 + \beta_W \cdot W_s + \beta_B \cdot B_s + \beta_L \cdot L_s \geq 0$, Male: $\beta_0 + \beta_W \cdot W_s + \beta_B \cdot B_s + \beta_L \cdot L_s < 0$, where intercept values (β_0) are −1.198 for *G. c. pratensis* and 1.533 for *G. c. tabida*. Scaled coefficients (i.e., scaled to unit length) permit judgement of relative importance of discriminating variables within subspecies.

greater accuracy and unison call posture is the most accurate method of determining sandhill crane gender in the field. However, gender classification errors of approximately 5%, at a cost of 20% of the KS-UG sample, may be tolerable when estimating demographic attributes of gender in high-powered situations, i.e., studies with large samples, small attribute variability, or appreciable attribute size or size difference.

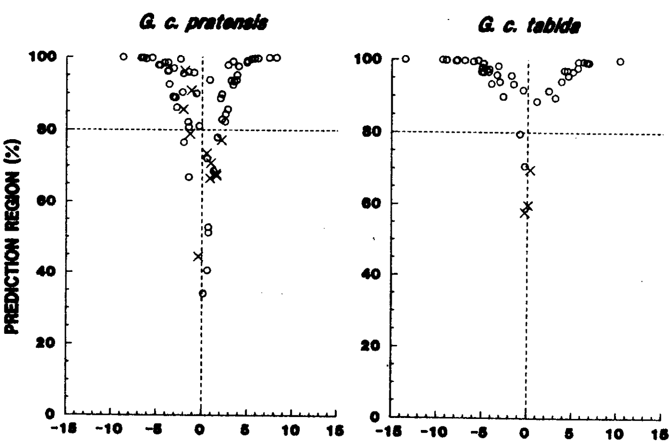

Fig. 2. Distance (prediction region size, %) to the farthest gender mean plotted against the jackknifed measurement discriminant score, by subspecies, for 114 known-gender sandhill cranes ≥18 months old caught in Florida, 1973–89. Cranes with positive (negative) discriminant scores were classified female (male) either correctly (o) or incorrectly (x). Misclassified observations typically fell within the 80% prediction region of both gender measurement means.

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